J Popul Econ (2011) 24:1385–1409
DOI 10.1007/s00148-010-0318-8

ORIGINAL PAPER

# Equilibrium unemployment and the duration of unemployment benefits

**Rafael Lalive · Jan C. van Ours · Josef Zweimüller**

Received: 8 June 2007 / Accepted: 26 February 2010 /
Published online: 29 April 2010


**Abstract** This paper uses microdata to evaluate the impact on the *steady-state unemployment rate* of an increase in maximum benefit duration. We evaluate a policy change in Austria that extended maximum benefit duration and use this policy change to estimate the causal impact of benefit duration on labor market flows. We find that the policy change leads to a significant increase in the steady-state unemployment rate and, surprisingly, most of this increase is due to an increase in the inflow into rather than the outflow from unemployment.

**Keywords** Benefit duration · Unemployment flows · Equilibrium unemployment

**JEL Classification** C41 · J64 · J65

*Responsible editor*: Christian Dustmann

R. Lalive
Faculty of Business and Economics, University of Lausanne, CEPR and IZA,
1015 Lausanne, Switzerland

J. C. van Ours
Department of Economics, CentER, Tilburg University, CEPR and IZA,
PO Box 90153, 5000 LE Tilburg, The Netherlands

J. Zweimüller (✉)
Institute for Empirical Research in Economics, University of Zurich, CEPR and IZA,
Mühlebachstrasse 86, 8008 Zurich, Switzerland
e-mail: zweim@iew.uzh.ch

## 1 Introduction

According to standard job search theory, more generous unemployment benefits increase the unemployment rate by reducing the search effort of workers thereby reducing the unemployment outflow rate. This prediction has been studied extensively in theoretical and empirical work and has proved to be empirically relevant and quantitatively important. The general finding from the empirical literature which will be discussed in more detail below is that it takes about 14 weeks of benefit duration to increase unemployment duration by one week.

The benefit system may affect unemployment not only via a reduced outflow from unemployment but also via a higher inflow into unemployment. One prominent argument, due to Mortensen and Pissarides (1994), holds that idiosyncratic shocks to workers' productivities let firms' optimal layoff rule depend on the wage rate—which in turn is affected by the prevailing unemployment benefit system. If the benefit system becomes more generous newly established jobs become unprofitable more quickly. As a result, a more generous benefit system will lead to an increase in the steady state flow from employment to unemployment. Alternatively, when workers' preferences (rather than their productivity) change randomly over time, a sufficiently negative shock may induce an employed worker to "quit" and collect benefits. More generous benefits will induce individuals to quit more easily raising the inflow into unemployment. As we discuss below, and in contrast to outflow effects, empirical evidence on the effects of benefit generosity on the unemployment inflow is much more scarce and far from conclusive.

The aim of this paper is to assess how the potential duration of unemployment benefits affects the equilibrium unemployment rate. Our main contribution is the analysis of the joint effects of benefit duration on the outflow from *and* the inflow into unemployment. This is different from the literature which has studied one of the two effects in isolation. The majority of previous studies has concentrated on the effects of the generosity of the benefit system on the probability that unemployed workers find regular jobs while a smaller literature has looked on the role of benefit rules on the probability to enter unemployment.

Understanding the inflow and outflow effects of the unemployment benefit system is crucial for labor market policy. First, the overall effect of a policy change remains unclear without a comprehensive understanding of both the inflow and the outflow channel. The risk is that policy makers may underestimate the implications of extended benefits for steady state labor market outcomes. Second, it is also crucial to understand the relative importance of the inflow and the outflow channel from a welfare point of view. Generous benefits that prolong unemployment spells can be problematic because long-term unemployment can cause skill depreciation. Skill depreciation is less of a concern when generous benefits mainly reduce job duration. As previous studies were typically concerned either with the inflow effect or with the outflow effect,

the relative size of these two effects remains unclear. The current study aims to shed light on their relative importance. As far as we know, this is the first paper that investigates the implications of the unemployment benefit system from a comprehensive perspective.[1]

Our analysis is based on a change in the Austrian unemployment insurance system that lead to a quasi-experimental situation allowing us to estimate benefit-duration effects on flows in and out of unemployment. In August 1989, the Austrian government made unemployment insurance more generous by increasing the maximum duration of unemployment benefits for certain groups of workers. Depending on age and previous work experience, the potential duration of regular benefits was raised from 30 to 52 weeks for one group, from 30 to 39 for a second group, and remained unchanged for a further group. We exploit this policy change and its differential treatment of these various groups of workers to assess the impact of benefit duration on unemployment inflows and outflows.

A particular advantage of our analysis is a very large and informative data set drawn from two sources: the Austrian unemployment register and the Austrian Social Security Data (ASSD). These data sources contain the universe of all employed and unemployed Austrian workers. We observe these worker over a period of four years, two years before the policy change, i.e. from August 1987 to July 1989; and two years after this policy change, from August 1989 to July 1991. A further advantage of our study concerns the fact that the period during which the policy change took place was quite stable from a macroeconomic perspective. This implies that our study is not subject to endogenous policy bias which arises when more generous unemployment insurance rules are implemented in anticipation of a deteriorating labor market. Such a policy bias has been found important in several recent studies (Card and Levine 2000; Lalive and Zweimüller 2004a). The absence of an endogenous policy bias, the large size and the low measurement error in our data set allow us to estimate the relevant policy parameters quite precisely.

Although we study both inflow and outflow effects of extended unemployment benefits the novelty of our paper is in the inflow analysis and above all in the comprehensive perspective. In previous work we analyzed outflow effects of the Austrian benefit system in detail. In Lalive et al. (2006) we show that the duration of unemployment is affected by two key parameters of unemployment insurance, the benefit replacement rate and the potential benefit duration. While the current study also addresses benefit duration effects, the overlap is limited. It is restricted to the relationship between

[1]There are cross-country studies that relate aggregate parameters of the unemployment insurance system—i.e. average replacement rate and average benefit duration—and other labor market institutions in various countries to the aggregate unemployment rates in these countries. See for an overview Layard and Nickell (1999).

potential benefit duration and outflow from unemployment.[2] In Lalive et al. (2006) we analyze the joint effect of potential benefit duration and changes in the replacement ratio. (In the present analysis we focus only on those workers for whom the replacement ratio remained unchanged). Moreover, the analysis is based on a more sophisticated methodology, a proportional hazard approach that allows us to investigate in detail the effect of observables on the exit rate from unemployment and the evolution of the exit rate over the elapsed duration. Because here we follow a more comprehensive approach in which we want to analyze inflow to and outflow from unemployment in a similar way we do not analyze the outflow processes in as much detail. In fact, as discussed in more detail below, we use straightforward logit analysis focusing on the probability to leave unemployment within a particular interval. The logit analysis is not as rich as the proportional hazard analysis but by and large generates similar results in terms of the way the extended potential benefit duration influence the overall outflow from unemployment. By performing a similar logit analysis for the probability to lose a job within a particular calendar time interval, we are able to capture equilibrium unemployment as derived from dynamics concerning both inflow and outflow.

Our findings with respect to the effect of the potential benefit duration (PBD) on the outflow from unemployment are in line with Lalive et al. (2006). The increase in PBD reduces the outflow a lot. The novel findings are twofold. First, the extension of the PBD also increases the inflow into unemployment. Our second finding is that the effect on the equilibrium unemployment rate due to increase in the inflow into unemployment is more important than the effect due to the decrease in the outflow from unemployment. Although the PBD extension makes it only a little bit more attractive for employed workers to become unemployed, there are many more employed workers than unemployed workers. This difference in the size of the two groups of workers causes the inflow effect to be larger.

The set-up of the paper is as follows. In Section 2 we review the relevant theoretical and empirical literature. Section 3 discusses the characteristics of the Austrian unemployment insurance system and briefly describes the Austrian labor market during the period when the change in maximum unemployment benefits was implemented. Section 4 presents the data we use in our analysis and discusses our empirical strategy. Section 5 presents parameter estimates

[2]Note that Lalive and Zweimüller (2004a, b) also use Austrian data to analyze how unemployment benefits affect the outflow from unemployment but these studies are based on information from Austrian regions with a dominant steel industry. In these regions, in 1988 an extended benefit program was introduced for workers aged 50 or older. The focus of both studies is on policy endogeneity, which indeed turns out to introduce a substantial bias in the parameter estimates. In Lalive et al. (2006) and the current paper to avoid policy endogeneity problems the analysis excludes the steel dominated regions.

and Section 6 uses our estimates to simulate the implied effects for the steady-state unemployment rate. Section 7 concludes.

## 2 How potential benefit duration affects unemployment

### 2.1 Theory

Denote by $\theta_{u,t}(x|T)$ the probability that an unemployed worker with personal characteristics $x$ finds a job in calendar time interval $t$ when $T$ is the maximum benefit duration (or potential benefit duration—PBD); and by $\theta_{e,t}(x|T)$ the probability that an employed worker with these characteristics loses his/her job in calendar time interval $t$. The steady state unemployment rate of the group of workers with characteristics $x$ is then

$$u^*(x|T) = \frac{\theta_e(x|T)}{\theta_e(x|T) + \theta_u(x|T)}. \tag{1}$$

Consider the effects of a change in the maximum benefit duration $T$ from the perspective of search theory. According to Mortensen (1977) expanding the duration of benefits has two opposite effects on the exit rate out of unemployment. First, the value of being unemployed increases so there is a disincentive effect that leads an unemployed worker to search less intensively. Second, the value of being employed also increases (because the value of being unemployed in the future has increased) which has a positive effect on the exit rate. For short-term unemployed the disincentive effect dominates, for unemployed near the point of benefit exhaustion (and beyond) the incentive effect dominates. Therefore, if there is an extension of benefit duration this will have a negative effect on the exit rate out of unemployment for short-term unemployed but it will have a positive effect on the exit rate for long-term unemployed. While the first effect has been found often in empirical research, evidence for the second effect is scarce (Fredriksson and Holmlund 2006).

The increase in the value of being unemployed through the extension of the potential benefit duration may also induce an increase in the inflow into unemployment. There are various reasons why this could be the case. For instance, the standard search and matching model with endogenous job destruction (Mortensen and Pissarides 1994, and Pissarides 2000, chapter 2) assumes that a worker's productivity on the job is subject to idiosyncratic shocks and firms require a minimum productivity level that prevents them from firing the worker and destroying the job. The firms' reservation productivity increases with more generous unemployment benefits, because more generous benefits push up wages requiring a higher average productivity on the job. Alternatively, assume a worker's disutility of labor (rather than his/her productivity) is subject to idiosyncratic shocks. In that case, more generous benefits will induce a worker to quit his/her job more easily. More generous unemployment benefits may also increase the take-up of unemployment benefits. Conditional on losing his/her job, a worker may be more inclined to apply for benefits.

Finally, it may be that the separation rate increases because a worker reduces his/her effort on the job and is more likely to be dismissed because he/she is less likely to take actions to prevent job loss.[3]

In conclusion, from a theoretical point of view, it is likely that $\partial\theta_u(x|T)/\partial T < 0$ and $\partial\theta_e(x|T)/\partial T > 0$. Therefore, an extension of the maximum benefit duration will increase the equilibrium unemployment rate:

$$\frac{\partial u^*(x|T)}{\partial T} > 0. \tag{2}$$

## 2.2 Empirical literature

Several US studies estimate the effects on the unemployment exit rate of variations in PBD that take place during recessions.[4] Early studies, including Moffitt and Nicholson (1982), Moffitt (1985), and Grossman (1989) find significantly negative incentive effects. Meyer (1990) and Katz and Meyer (1990) show that the exit rate from unemployment rises sharply just before benefits are exhausted. Such spikes are absent for nonrecipients. More recent work by Addison and Portugal (2004) confirms these findings.[5]

A common objection against these studies is policy endogeneity. Benefits are typically extended in anticipation of a worse labor market for the eligible workers. Card and Levine (2000) exploit a variation in benefit duration in New Jersey that occurred independently of labor market condition and show that policy bias is substantial. Lalive and Zweimüller (2004b) find similar evidence for Austria.

Evidence on the effect of PBD in European studies is mixed. Hunt (1995) finds substantial disincentive effects of extended benefit entitlement periods for Germany. Carling et al. (1996) find a big increase in the outflow from unemployment to labor market programs whereas the increase in the exit rate to employment is substantially smaller. Puhani (2000) finds that reductions in PBD in Poland did not have a significant effect on the duration of unemployment whereas Adamchik (1999) finds a strong increase in reemployment probabilities around benefit expiration. Roed and Zhang (2003) find for Norwegian unemployed that the exit rate out of unemployment increases sharply in the months just prior to benefit exhaustion where the effect is larger for females than for males. Winter-Ebmer (1998) and Lalive

[3]Note, however, that according to Fredriksson and Holmlund (2006) there is not much empirical evidence in support of such an effect.

[4]Fredriksson and Holmlund (2003) give a recent overview of empirical research related to incentives in unemployment insurance. See Green and Riddell (1997), and Ham and Rea (1987) for studies that focus on Canada.

[5]Note that there is no theoretical explanation for the existence of end-of-benefit spikes. It could be that the spikes have to do with strategic timing of the job starting date, i.e. workers have already found a job but they postpone starting to work until their benefits are close to expiration. Card and Levine (2000) point at the possibility that there is an implicit contract between the unemployed worker and his previous employer to be rehired just before benefit expire.

and Zweimüller (2004b) show that extending the potential duration of benefits had significant disincentive effects in Austria. Van Ours and Vodopivec (2006) studying PBD reductions in Slovenia find both strong effects on the exit rate out of unemployment and substantial spikes around benefit exhaustion.

Empirical studies on the unemployment inflow effect of a lengthening of the maximum benefit duration are more rare. Most of these studies focus on requirements concerning entrance into unemployment insurance. Christofides and McKenna (1995, 1996) for example find a clear relationship between entrance requirements of Canadian unemployment insurance and employment durations. The exit rate from employment to unemployment increases substantially as soon as the workers satisfy the number of weeks worked in order to qualify for UI benefits. Anderson and Meyer (1997) investigate the take up rate of unemployment benefit insurance of workers separating from their employer. They find that both the level and the maximum duration of benefits have a significant positive effect. Green and Riddell (1997) study the effect of changes in entrance requirements on the inflow into Canadian unemployment finding that changes in these requirements have a significant impact on employment durations. They also find that many employment spells that just qualify under the old system are extended to just qualify under the new system. And they find that all of the response is in layoffs, not quits, which suggests that employers play an important role in the adjustment of employment durations. Green and Sargent (1998) analyze Canadian data and also find evidence of concentrations of job spell durations at the entrance requirement point and at the point at which individuals have qualified for the maximum possible weeks of benefit receipts. Winter-Ebmer (2003) finds strong inflow effects of the Austrian regional extended benefit program which granted very long benefits for older workers in certain regions.[6] These results are in line with those of Lalive and Zweimüller (2004a) who also find significant inflow effects which were particularly strong immediately before this program was abolished.

## 3 Institutional background

Like in a number of other countries the Austrian unemployment insurance system is characterized by a limited period over which unemployed individuals can draw 'regular' unemployment benefits. Unemployment benefits depend on previous earnings and, compared to other European countries, the replacement ratio (benefits relative to *gross* monthly earnings) is rather low. In 1990, the replacement ratio was 40.4% for the median income earner; 48.2% for a low-wage worker who earned half the median; and 29.6% for a high-wage worker earning twice the median. On top, family allowances

[6]The regional extended benefit program was implemented in 1987 and ended in 1993 and was directed to a subset of Austrian regions. (See Winter-Ebmer 1998, 2003 and Lalive and Zweimüller 2004a, b). The policy change analyzed here applies to workers in all other regions and excludes regions that were subject to the regional extended benefit program.

are paid. Unemployment benefit payments are not taxed and not means-tested. Voluntary quitters and workers discharged for misconduct cannot claim benefits until a waiting period of 4 weeks has passed. Unemployment benefit recipients are expected to search actively for a new job that should be within the scope of the claimant's qualifications, at least during the first months of the unemployment spell. Non-compliance with the eligibility rules is subject to benefit sanctions that can lead to the withdrawal of benefits for up to 4 weeks.

Once the period of regular unemployment benefits has expired, individuals can apply for "transfer payments for those in need".[7] As the name indicates, these transfers are means-tested and the job seeker is considered eligible only if she or he is in trouble. These payments depend on the income and wealth situation of other family members and close relatives and may, in principle, last for an indefinite time period. These transfers are granted for successive periods of 39 weeks after which eligibility requirements are recurrently checked. The post-unemployment benefits transfers are lower than unemployment benefits and can at most be 92% of unemployment benefits. In 1990, the median post-unemployment benefits transfer payment was about 70% of the median unemployment benefits. Note, however, that individuals who are eligible for such transfers may not be comparable to individuals who collect unemployment benefits because not all individuals who exhaust unemployment benefits pass the means test. The majority of the unemployed (59%) received unemployment benefits, whereas 26% received post-unemployment benefit transfers. In sum, the Austrian unemployment insurance system is less generous than many other continental European systems and closer to the U.S. system.[8]

Before August 1989, an unemployed person could draw regular unemployment benefits for a maximum period of 30 weeks provided that he or she had paid unemployment insurance contributions for at least 156 weeks within the last 5 years.[9] In August 1989 the potential duration of unemployment benefit payments became dependent not only on previous experience but also on age at the beginning of the unemployment spell. Benefit duration for the age group 40–49 was increased to 39 weeks if the unemployed had 312 weeks of employment within the last 10 years prior to the current spell. For the age group 50 and older, unemployment benefit duration was increased to 52 weeks if the unemployed had been employed for at least 468 weeks within the last 15 years. Austrian policy makers introduced age delineation for two reasons. First, as age is the strongest predictor of long-term unemployment, policy

[7]This so-called "Notstandshilfe" implies that job seekers who do not meet benefit eligibility criteria can apply at the beginning of their spell.

[8]See Nickell and Layard (1999). It is interesting to note that the incidence of long-term unemployment in Austria is closer to U.S. figures than to those of other European countries. In 1995, when our sample period ends, 17.4% of the unemployment stock were spells with an elapsed duration of 12 months or more. This compares to 9.7% for the U.S. and to 45.6% for France, 48.3% for Germany, and 62.7% for Italy (OECD 1999).

[9]UB duration was 20 weeks for job-seekers who did not meet this requirement. This paper focuses on individuals who were entitled to at least 30 weeks of benefits.

makers wanted to improve protection for older workers by granting payment of regular unemployment benefits for a longer time period. Second, about one year prior to the August 1989 policy change, the *Krisenregionsregelung* introduced very strong age delineation at age 50 years. Age delineation was therefore adopted in August 1989 to be consistent with earlier modification of unemployment benefit rules.

## 4 Data and empirical strategy

To assess the impact of increasing benefit duration on unemployment outflow and inflow rates, we use longitudinal individual data from two different sources: (i) the *Austrian Social Security Database* (ASSD), which contains detailed information on the individuals' employment, unemployment and earnings history since the year 1972, and some information on the employer, like region and industry affiliation; and (ii) the *Austrian unemployment register*, from which we get information on the relevant socioeconomic characteristics.

From these data we drew two samples, a "before-policy" sample and an "after-policy" sample, as follows. The sample was constructed such that all individuals fulfill the experience requirement required for extended eligibility and that control individuals not too different in age from the treatment group. Thus, for both samples we selected individuals who were at least 35 years and at most 54 years old. Furthermore, we included only individuals with a continuous work history. To be included in the sample, an individual had to have a job for at least 6 out of the last 10 years and for at least 9 out of the last 15 years. Hence all individuals in our sample satisfy the work experience criteria for eligibility to extended benefit duration (see above). Our additional sample restrictions were motivated by concerns with events that may confound the estimates of the effects of PBD on inflow and outflow. We excluded all individuals living in regions subject to the regional extended benefit program because these regions were covered by different extended benefit rules for workers older than 50 years.[10] Furthermore, we considered only workers with previous income above Austrian Shilling 12,610 (Euros 916) because workers below this threshold experienced an increase in benefits levels in 1989 a policy change that is beyond the scope of this paper. The sample also excludes workers previously employe in the construction and tourism industries as unemployment durations in these groups are mainly driven by seasonal factors.

To make the sample as clean as possible we concentrate our analysis on "attached workers", i.e. workers who are either employed or unemployed during the first observation year. The reason is that our focus is on flows between employment and unemployment rather than on flow between employment

[10]This so-called *Krisenregionsregelung* applied to about 15% of all observations. In these crises-ridden regions even more generous unemployed insurance policies were implemented between 1988 and 1993. For empirical analyzes of these programs, see Winter-Ebmer (1998, 2003) and Lalive and Zweimüller (2004a, b).

(unemployment) and out-of-labor force. While this is potentially interesting our data set is not informative whether individuals have access to programs other than unemployment insurance in case they drop out of the work force. Concentrating on attached workers has the advantage that flows in and out of non-employment are less important. Our empirical analysis is based on a comparison of two samples. The "before-policy" sample contains individuals who were either (i) employed as white or blue collar workers or who were (ii) unemployed at one of the four mid-quarter baseline dates (10th February, 10th May, 10th August, 10th October) in the year 1986. We then follow these individuals up until quarter IV.1988. The "after-policy" sample contains all individuals who were either unemployed or employed as blue- or white-collar workers at each mid-quarter baseline date in the year 1990. We then follow these individuals up until quarter IV.1992. Note that this design allows individuals in the before-policy (after-policy) sample to be *out of labor force* only in the years 1987 and 1988 (1991 and 1992). Hence this restriction reinforces our focus on attached workers whose unemployment rates are below the actual unemployment rates which include workers with some distance from the labor market. Note further, that we do not consider observations for the year 1989. This procedure minimizes potential biases resulting from anticipation effects that may arise due to behavioral changes of individuals that were unemployed under initial policy rules but were anticipating that rules will become more generous.

Table 1 compares the characteristics of the two groups. There are basically two major differences between the two groups. First, we see that after the policy change, somewhat more than a quarter (half) of the sample is eligible to additional 22 (9) weeks of potential benefits duration. While average age in the before-policy sample is only slightly younger (by 0.4 years) than the after-policy sample, the distribution across relevant age groups is more strongly affected. Second, we see that the after-policy sample has a higher fraction of females.[11] Otherwise, the differences between samples are minor. Real earnings are slightly higher in the after-policy sample. Years of work experience within the last 15 years ("Experience") and the duration of the current job ("Tenure"; for the non-employed: tenure in the last job) is slightly higher in the after policy sample. Moreover, the number of white collar workers and the industry distributions of the two samples are very similar.

Our analysis of the impact of the maximum duration of benefits on the steady-state unemployment rate is based on an analysis of individual transition probabilities to and from unemployment. To assess the effect of the maximum

[11]The higher fraction of ages 50+ is because the big birth cohorts of 1940–1942 are in the age group 40–49 in the before-policy sample whereas they are in the age group 50+ in the after-policy sample. The higher fraction of females in the after-policy sample is most likely due to the fact that the cohorts that are in the after-policy but not in the before-policy sample have a high labor force participation and are relatively large (vintages in the mid 1950s). In contrast, the cohorts that are in the before-policy sample but not in the after-policy sample (vintages of the early 1930s) do have a low labor force participation and are comparably small.

**Table 1** Descriptive statistics

| | Unemployment outflow | | | | Unemployment inflow | | | |
|---|---|---|---|---|---|---|---|---|
| | Before-policy sample | | After-policy sample | | Before-policy sample | | After-policy sample | |
| | Mean | Std. dev. | Mean | Std. dev. | Mean | Std. dev. | Mean | Std. dev. |
| Eligible for 52 weeks | 0.000 | 0.000 | 0.382 | 0.486 | 0.000 | 0.000 | 0.308 | 0.462 |
| Eligible for 39 weeks | 0.000 | 0.000 | 0.492 | 0.500 | 0.000 | 0.000 | 0.539 | 0.498 |
| Eligible for 30 weeks | 1.000 | 0.000 | 0.126 | 0.332 | 1.000 | 0.000 | 0.153 | 0.360 |
| Duration of unemployment (weeks) | 41.630 | 63.845 | 48.977 | 76.780 | 0.000 | 0.000 | 0.000 | 0.000 |
| Age 35–39 | 0.181 | 0.385 | 0.126 | 0.332 | 0.173 | 0.378 | 0.153 | 0.360 |
| Age 40–49 | 0.562 | 0.496 | 0.492 | 0.500 | 0.580 | 0.494 | 0.539 | 0.498 |
| Age 50+ | 0.257 | 0.437 | 0.382 | 0.486 | 0.247 | 0.431 | 0.308 | 0.462 |
| Age | 45.383 | 5.500 | 46.768 | 5.526 | 45.474 | 5.457 | 45.941 | 5.416 |
| After-policy | 0.000 | 0.000 | 1.000 | 0.000 | 0.000 | 0.000 | 1.000 | 0.000 |
| 1st quarter | 0.351 | 0.477 | 0.301 | 0.459 | 0.250 | 0.433 | 0.250 | 0.433 |
| 2nd quarter | 0.220 | 0.414 | 0.219 | 0.414 | 0.251 | 0.434 | 0.251 | 0.434 |
| 3rd quarter | 0.212 | 0.408 | 0.229 | 0.420 | 0.250 | 0.433 | 0.250 | 0.433 |
| 4th quarter | 0.218 | 0.413 | 0.251 | 0.434 | 0.248 | 0.432 | 0.248 | 0.432 |
| (log) wage | 6.425 | 0.276 | 6.440 | 0.316 | 6.573 | 0.280 | 6.598 | 0.300 |
| Experience[a] (years) | 13.143 | 1.487 | 13.102 | 1.520 | 14.293 | 1.241 | 14.252 | 1.293 |
| Tenure (years) | 5.209 | 5.514 | 6.062 | 6.635 | 10.267 | 5.522 | 11.062 | 6.954 |
| White collar | 0.151 | 0.358 | 0.158 | 0.364 | 0.596 | 0.491 | 0.606 | 0.489 |
| Manufacturing | 0.449 | 0.497 | 0.423 | 0.494 | 0.415 | 0.493 | 0.384 | 0.486 |
| Women | 0.349 | 0.477 | 0.436 | 0.496 | 0.333 | 0.471 | 0.378 | 0.485 |

[a]Work experience during last 15 years

benefit duration on these transition probabilities we use a simple difference-in-differences estimator in the context of a logit model for quarterly transition probabilities for observation $i$ in quarter $t$

$$\theta^*_{yit} = \delta_{y1} ELIG52_{it} + \delta_{y2} ELIG39_{it} + \gamma_y A_{it} + x_{it}\beta_y + \varepsilon_{yit} \quad (3)$$
$$\theta^*_{yit} \geq 0 \text{ if } \theta_{yit} = 1 \quad \text{and} \quad \theta^*_{yit} < 0 \text{ if } \theta_{yit} = 0$$

where $y$ is a subscript indicating whether the transition concerns outflow from unemployment ($y = u$) or inflow into unemployment ($y = e$). The variables $ELIG52_{it}$ and $ELIG39_{it}$ are indicator variables that take value 1 when observation $i$ is eligible for at most 52 or at most 39 benefit weeks, respectively.[12] Furthermore, $\delta_{y1}$ and $\delta_{y2}$ are the corresponding differences-in-differences estimators, the dummy variable $A_{it}$ indicates the after-policy period and $\gamma_y$ measures the calender time effect on transition $y$ that is irrespective of observation $i$'s eligibility status. Finally, $x_{it}$ is a vector in individual characteristics, $\beta_y$ is a vector of parameters that estimate the impact of these characteristics on transition $y$,[13] and the error term $\varepsilon_{yit}$, capturing unobservable heterogeneity, are assumed to be standard normally distributed.[14]

Obviously, whether the difference-in-differences estimator identifies the causal effect of the increase in benefit duration on the unemployment risk hinges upon whether or not the policy change was exogenous.[15] There are two reasons why policy endogeneity is most likely of minor importance in the present context. The first reason is that the economy was doing badly before the policy change (in the years 1987 and 1988). After the policy change (in the years 1989, 1990, 1991) the economy was in a boom. To the extent that all age groups were benefitting from this situation, policy endogeneity is not an issue. Second, one reason for the implementation of the policy may have been equity concerns. In 1988, the Austrian government implemented a very generous program that was targeted towards older steel workers in crises ridden steel regions. This 'Austrian regional extended benefit program' granted 4 years of unemployment benefits to eligible older workers in crisis-ridden steel regions. Hence political pressure to treat older unemployed workers in non-eligible regions more generously was one reason for changing the benefit rules. To the extent that such equity concerns were the reason for the policy change, the increase in benefit duration can be regarded as exogenous with respect to labor market outcomes of the eligible individuals in our sample.

[12] All observations in our samples for which both $T39_i = 0$ and $T52_i = 0$ are eligible for at most 30 weeks of benefits.

[13] The vector of individual characteristics includes the individual's age, age dummies, dummies for the inflow quarter, log daily wage, experience, tenure, broad occupation (blue/white collar), sex, and industry (manufacturing, construction/tourism, other industries).

[14] The analysis below will be undertaken also for more flexible specifications of age and calendar time, and will be estimated for various subgroups to assess the robustness of the results.

[15] If policy was implemented because policy makers became concerned with worse labor market prospects for older individuals there would be policy endogeneity.

An important assumption of the diff-in-diff strategy is that there are no differential trends in unemployment between treated and control groups. In our analysis below we will use the age group 35–39 as the benchmark for changes in unemployment inflows and outflows in the absence of the PBD extension. To check whether this age group is a good benchmark for the changes in unemployment Fig. 1 provides information on trends in unemployment during the pre-reform period. Note that Fig. 1 is based on the official unemployment rate thus covering both attached workers (as our sample does) as well as unattached workers leading to higher unemployment rates than in our sample. As Fig. 1 shows clearly that the unemployment rates of the age groups 35–39, 40–49 and 50–54 are similar and do not show any substantial differences in the trend over the pre-reform period.

Our analysis concentrates on the period 1987–1992. In 1987 the economy was at the end of a recession and started to improve. Real GDP growth was 1.7% in 1987 and then started to grow to as much as 4.7% in 1990. The favorable situation of the business cycle led to strong employment growth throughout the period under consideration. The unemployment rate was rather stable due to an increase in labor supply (immigration and rising female labor force participation). Aggregate flows into and out of unemployment did not dramatically change during the period under consideration. The aggregate quarterly unemployment inflow rate (new unemployment spells that started in given quarter relative to the total stock of employment and out-of-labor-force) was fluctuating around 2.75% and the average duration of unemployment (spells completed during respective year) was roughly stable at somewhat less than 4 months. The average unemployment rate during the post-treatment period 1989–1991 was as high as during the pre-treatment period 1987–1988. Furthermore, employment growth during the treatment period was even somewhat stronger than before.

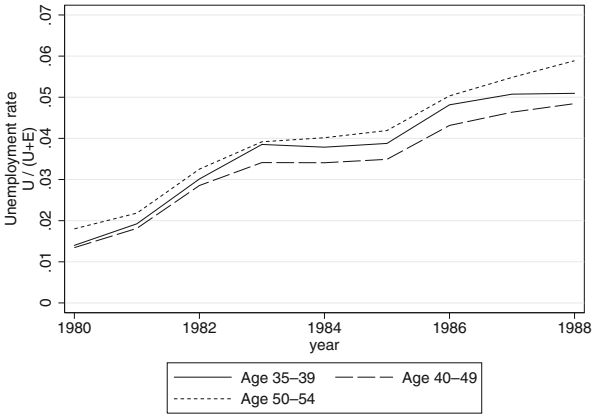


**Fig. 1** Pre-reform trends in unemployment rates, by age groups

Notes: U: # of unemlpoyed workers, E: # of employed workers.
Source: Own calculations based on ASSD.

It is worth noting that the improving labor market is favorable in terms of our empirical strategy. This is because it is unlikely that comparing labor market experiences of older workers before and after the policy change are driven by a deteriorating labor market. Notice, however, that workers aged 35–39 could be affected by business cycle fluctuations in a different way than older workers aged 40–49 or 50–54 invalidating the assumption of diff-in-diff analysis. As a crude check whether this concern is relevant we ran separate OLS regressions with the age-specific unemployment rate as the dependent variable and the aggregate unemployment rate as a regressor. For all three age groups, the coefficients are not significantly different from unity and point estimates are almost identical in these regressions. Hence business-cycle reactions do not seem to differ strongly across age-groups.

## 5 Empirical estimates

We proceed in two steps. We first show the regression results of our basic statistical model for the unemployment flows, separately for the unemployment outflow and the unemployment inflow. We next check (i) whether our estimated effects of benefit duration extension are robust to a more flexible specification of the age and calendar time variables; (ii) how the estimated effects of benefit duration are robust once the duration of the current state (elapsed duration of unemployment in the outflow equation; and tenure on the current job in the inflow equation); and (iii) how the estimated effects differ across various population subgroups. Using our outflow- and inflow-estimates, we can then discuss the question of interest: How do the changes in maximum benefit duration affect the steady-state unemployment rate?

### 5.1 Unemployment outflow and inflow

Table 2 shows the results of equation (3) both for the unemployment outflow (columns 1 and 2) and the unemployment inflow (columns 3 and 4). (Notice that the coefficients displayed in the Table are marginal effects evaluated at the sample means of the right-hand-side variables). The logit estimation of column 1 includes all 29,786 unemployment cases observed in one of the eight quarterly baseline dates. Similarly, the estimation of column 2 is based on the 1,245,337 employment cases observed in our sample.

The diff-in-diff estimators are in line with the theoretical predictions. Eligibility to longer benefits reduces the outflow rate from unemployment (column 1) and increases the inflow probability into unemployment (column 2). All coefficients have the expected sign. The effect of increasing PBD by 22 weeks is particularly strong, both in the outflow and in the inflow equation. The effect of increasing PBD by 9 weeks is weaker and statistically significant only in the inflow equation.

**Table 2** Logit results on unemployment flows

| Dependent variable | Unemployment outflow | | Unemployment inflow | |
|---|---|---|---|---|
| Mean of dep. variable | 0.445 | | 0.011 | |
| | (1) | (2) | (3) | (4) |
| Eligible for 52 weeks | −0.090 | −0.041 | 0.002 | 0.003 |
| | (4.30)** | (2.43)* | (4.19)** | (5.77)** |
| Eligible for 39 weeks | −0.024 | −0.018 | 0.001 | 0.001 |
| | (1.22) | (1.15) | (2.25)* | (3.78)** |
| Duration of unemployment (weeks) | | −0.008 | | |
| | | (24.30)** | | |
| Tenure (years) | | | | −0.001 |
| | | | | (74.23)** |
| After-policy | −0.208 | −0.141 | −0.000 | −0.001 |
| | (9.23)** | (6.80)** | (0.64) | (2.57)* |
| (log) wage | −0.150 | −0.054 | −0.008 | −0.004 |
| | (11.37)** | (5.17)** | (23.60)** | (13.38)** |
| Experience (years) | −0.013 | −0.024 | −0.003 | −0.001 |
| | (5.13)** | (12.06)** | (59.81)** | (36.37)** |
| White collar | 0.089 | 0.044 | −0.004 | −0.002 |
| | (9.63)** | (4.95)** | (16.06)** | (13.16)** |
| Manufacturing | −0.022 | −0.016 | 0.001 | 0.001 |
| | (3.03)** | (2.85)** | (3.36)** | (5.74)** |
| Women | −0.096 | −0.084 | −0.003 | −0.001 |
| | (12.59)** | (13.87)** | (16.61)** | (5.07)** |
| Calendar time dummies (for each quarter) | Yes | Yes | Yes | Yes |
| Age dummies (for each year) | Yes | Yes | Yes | Yes |
| Observations | 29786 | 29786 | 1245337 | 1245337 |

Note: Marginal effects, absolute value of robust $z$ statistics in parentheses, clustered at the individual level

*Significant at 5%; **significant at 1%

The coefficient of column 1 in Table 2 indicates that the probability that an individual leaves the unemployment register within the next quarter is reduced by 9 percentage points for individuals that become eligible to a maximum of 52 benefit weeks instead of the 30 weeks before the policy change. Similarly, the point estimate of the increase from 30 to 39 weeks of maximum benefit duration indicates a (statistically insignificant) reduction of 2.4 percentage points. Column 2 in Table 2 checks how these effects change once we control for the elapsed duration of the individual's current unemployment spell. It turns out that the effects still significantly negative when becoming eligible to 52 benefit weeks albeit the estimated impact reduces to 4.1 percentage points whereas becoming eligible to 39 weeks leads only to a marginal change in the point estimate.

We find a strong impact of extending the maximum benefit duration on the unemployment inflow. The effects are not only highly significant, they are also quantitatively substantial. Column 3 of Table 2 suggests that, as a result of

an increase in maximum benefit duration from 30 to 52 weeks, the quarterly inflow rate increases by 0.2 percentage points. While this looks like a small number we have to keep in mind that the average quarterly unemployment inflow rate is also a small number which amounts to roughly 1% per quarter. Taken together, this results in a quantitatively substantial increase in the inflow rate of 10%. Similarly, the extension of maximum benefit duration from 30 to 39 weeks, leads to an increase is associated with a 0.1 percentage point increase or a 10% increase in the unemployment inflow rate. Column 4 includes the duration of the current job ("tenure") as an additional regressor in the estimated equation. It turns out that controlling for tenure increases the estimated point effects for workers who become eligible for 52 benefit weeks the inflow rate increases by .3 percentage points whereas the estimated effect for workers with 39 weeks remains roughly constant.[16]

Table 2 also displays the coefficient for the remaining control variables included in the regressions. We find that *high-wage* workers have substantially lower turnover: they do not only have a lower risk of unemployment but also a lower chance to exit unemployment have also lower chances to exit unemployment. Similar effects are found with respect to previous *work experience*. In contrast, *white-collar* workers have a higher unemployment exit rate and a lower unemployment entry rate. Results also show significant differences between *industries* with manufacturing workers showing lower turnover than other industries (mostly services). Finally, we find that *women* have significantly worse chances than men to exit unemployment. Furthermore, there are significant gender differences in the risk of unemployment.

To investigate further to which extent the duration in the current state may have an impact on the above results, Table 3 performs a number of further robustness checks. The upper and lower panels of this table refer to outflow and inflow regressions, respectively. For ease of comparison, column 1 in Table 3 repeats the baseline results obtained in columns 2 and 4 of Table 2. In column 2, we allow for interaction effects of duration with both calendar time and age dummies. This does not have any major impact on the estimated benefit duration effects. For individuals eligible to the extension to 52 weeks we find a slightly larger negative impact on the outflow and a somewhat lower impact on the unemployment inflow. For individuals eligible to the extension to 39 weeks, the coefficients remain basically unchanged. In the last two columns of Table 3 we split the sample into short and long durations (below and above 15 weeks of elapsed unemployment duration in the outflow sample; below and above 10 years of tenure in the inflow sample). It turns

[16]With respect to the effect of PBD on the unemployment outflow, our results are in line with the estimates in Lalive et al. (2006) who find that the increase in PBD from 30 to 52 weeks lead to an increase in the expected duration of unemployment of 12.3% and who find a very small effect of the increase in PBD from 30 to 39 weeks. Our results are also similar to previous estimate to Winter-Ebmer (2003) who finds substantial effects of PBD on the unemployment inflow for a different policy change in Austria, which extended PBD for older worker in certain regions.

**Table 3** Splitting the sample by duration of current state, logit estimates

| Unemployment outflow | Whole sample (1) | Whole sample (2) | Duration ≤15w (3) | Duration >15w (4) |
|---|---|---|---|---|
| Eligible for 52 weeks | −0.041 | −0.055 | −0.027 | −0.043 |
| | (2.43)* | (3.27)** | (0.85) | (3.01)** |
| Eligible for 39 weeks | −0.018 | −0.016 | −0.027 | −0.021 |
| | (1.15) | (1.03) | (0.96) | (1.54) |
| Duration of unemployment (weeks) | −0.008 | −0.005 | 0.008 | −0.004 |
| | (24.30)** | (2.95)** | (6.31)** | (17.28)** |
| Calendar time dummies (for each quarter) | Yes | Yes | Yes | Yes |
| Age dummies (for each year) | Yes | Yes | Yes | Yes |
| Interaction of duration with age and calendar time dummies | No | Yes | No | No |
| **Unemployment inflow** | **Whole sample (1)** | **Whole sample (2)** | **Tenure ≤10y (3)** | **Tenure >10y (4)** |
| Eligible for 52 weeks | 0.003 | 0.002 | 0.003 | 0.002 |
| | (5.77)** | (4.66)** | (3.31)** | (3.39)** |
| Eligible for 39 weeks | 0.001 | 0.001 | 0.002 | 0.000 |
| | (3.78)** | (3.47)** | (3.73)** | (0.26) |
| Tenure (years) | −0.001 | −0.001 | −0.003 | −0.000 |
| | (74.23)** | (5.90)** | (67.76)** | (4.71)** |
| Calendar time dummies (for each quarter) | Yes | Yes | Yes | Yes |
| Age dummies (for each year) | Yes | Yes | Yes | Yes |
| Interaction of tenure with age and calendar time dummies | No | Yes | No | No |

Note: Marginal effects, absolute value of robust $z$ statistics in parentheses, clustered at the individual level

* Significant at 5%; ** significant at 1%

out that outflow effects show up predominantly longer unemployment spells when benefits are extended to 52 weeks. While the remaining coefficients show the expected negative sign, they turn out statistically insignificant. In contrast, unemployment inflow effects are somewhat stronger for workers with low tenure and become insignificant for long-tenure workers eligible to 39 benefit weeks. In sum, while we find differential impacts of workers with different durations in their current state, the basic picture of Table 2 remains. In general inflow effects seem to be stronger than outflow effects.

As a further test for the robustness of our results, we look at the effects of the change in maximum benefit duration once we split our sample into various subgroups (Table 4). All regressions are based on the specification we used in Table 2, column 2 (outflow) and column 4 (inflow). The basic message of Table 4 is that increasing the maximum benefit duration both from 30 to 52 weeks and from 30 to 39 weeks has a significant and quantitatively important impact on the unemployment inflow. In contrast, the unemployment outflow coefficients are less important and the point estimates mostly insignificant.

**Table 4** Diff-in-diff logit marginal effects of PBD-effects: various subgroups

| | Unemployment outflow | | | Unemployment inflow | | |
|---|---|---|---|---|---|---|
| | Eligible for 52 weeks | Eligible for 39 weeks | Duration (weeks) | Eligible for 52 weeks | Eligible for 39 weeks | Tenure (years) |
| Whole sample | −0.041 | −0.018 | −0.008 | 0.003 | 0.001 | −0.001 |
| | (2.426)* | (1.150) | (24.299)** | (5.768)** | (3.780)** | (74.230)** |
| Women | −0.025 | 0.007 | −0.006 | 0.005 | 0.001 | −0.001 |
| | (0.896) | (0.285) | (10.760)** | (4.014)** | (1.725) | (35.070)** |
| Men | 0.020 | 0.025 | −0.010 | 0.002 | 0.001 | −0.001 |
| | (0.760) | (1.103) | (17.694)** | (3.880)** | (3.744)** | (57.779)** |
| Blue collar | −0.006 | 0.012 | −0.010 | 0.004 | 0.002 | −0.001 |
| | (0.185) | (0.418) | (12.194)** | (4.630)** | (3.394)** | (45.717)** |
| White collar | 0.019 | 0.011 | −0.006 | 0.003 | 0.001 | −0.001 |
| | (0.838) | (0.537) | (16.854)** | (4.595)** | (1.954) | (42.286)** |
| Low wage | −0.015 | 0.019 | −0.009 | 0.005 | 0.002 | −0.001 |
| | (0.604) | (0.847) | (16.364)** | (4.611)** | (2.705)** | (55.275)** |
| High wage | −0.002 | −0.018 | −0.008 | 0.002 | 0.001 | −0.001 |
| | (0.068) | (0.807) | (16.245)** | (3.871)** | (2.403)* | (52.639)** |

Note: Marginal effects, absolute value of robust $z$ statistics in parentheses, clustered at the individual level

* Significant at 5%; ** significant at 1%

## 6 Benefit duration and equilibrium unemployment

Using the parameter estimates of the inflow and outflow probabilities we consider how the maximum benefit duration affects equilibrium unemployment. Our thought experiment is the following. Let us take our estimates of the increase in PBD at face value and consider a steady-state situation in which the inflow into and the outflow from unemployment are identical. Which unemployment rate is implied by the system before the policy change as compared to the system after the change. Ignoring effects of personal characteristics $x$ we have

$$u^*(T) = \frac{\hat{\theta}_e(T)}{\hat{\theta}_e(T) + \hat{\theta}_u(T)} \tag{4}$$

The policy changes we are analyzing are discrete, and amount to a substantial increase in maximum benefit duration for the concerned groups. In order to assess the effect of the change in benefit duration on equilibrium unemployment, we perform a comparative static analysis. If $T_1$ and $T_2$ are the maximum benefit durations before and after the policy change, the change in equilibrium unemployment equals

$$\Delta u^* = u^*(T_2) - u^*(T_1) \tag{5}$$

Furthermore, it is straightforward to decompose this change into (i) a change due to a lower outflow from unemployment, (ii) a change due to a higher inflow into unemployment, and (iii) to an interaction effect involving higher-order terms

$$\Delta u^* = \Delta u^*(out) + \Delta u^*(in) + \text{ interaction effect}$$

where the inflow- and outflow-effects are given by

$$\Delta u^*(out) = \frac{\hat{\theta}_e(T_1)}{\hat{\theta}_e(T_1) + \hat{\theta}_u(T_2)} - \frac{\hat{\theta}_e(T_1)}{\hat{\theta}_e(T_1) + \hat{\theta}_u(T_1)}$$

$$\Delta u^*(in) = \frac{\hat{\theta}_e(T_2)}{\hat{\theta}_e(T_2) + \hat{\theta}_u(T_1)} - \frac{\hat{\theta}_e(T_1)}{\hat{\theta}_e(T_1) + \hat{\theta}_u(T_1)}$$

Notice that, just like in any other decomposition analysis there is an "interaction effect" (or unexplained residual). When we calculate $\Delta u^*(out)$, we change the outflow rate from the pre-policy level $\hat{\theta}_u(T_1)$ to the post-policy level $\hat{\theta}_u(T_2)$, keeping the inflow rate $\hat{\theta}_e(T_1)$ at its pre-policy level. Similarly, $\Delta u^*(in)$ calculates the hypothetical effect on equilibrium unemployment of an isolated change in the inflow rate keeping the outflow rate at its pre-policy level. Since equilibrium unemployment is non-linearly related to inflow and outflow rates, $\Delta u^*(in)$ and $\Delta u^*(out)$ do not sum up to $\Delta u^*$. The residual is captured by the interaction effect.

**Table 5** Effects of PBD increase in inflow, outflow and unemployment population ratio

| | Quarterly outflow | Quarterly inflow | Interaction | Implied steady-state unemployment ratio (%) |
|---|---|---|---|---|
| PBD change 30 to 52 weeks | | | | |
| Before policy-change | .4012 | .0098 | | 2.38 |
| After policy-change | .3680 | .0142 | | 3.72 |
| Implied increase in u* (p.p.) | .21 | 1.04 | .09 | 1.34 |
| (percentage due to ...) | (15.6%) | (77.8%) | (6.6%) | (100.0%) |
| Implied increase in u* per additional PBD week (p.p.) | 0.010 | 0.047 | 0.004 | 0.061 |
| PBD change 30 to 39 weeks | | | | |
| Before policy-change | .4649 | .0098 | | 2.07 |
| After policy-change | .4491 | .0121 | | 2.62 |
| Implied increase in u* (p.p.) | .07 | .46 | .02 | .55 |
| (percentage due to ...) | (13.1%) | (84.1%) | (2.8%) | (100.0%) |
| Implied increase in u* per additional PBD week | 0.008 | 0.051 | 0.002 | 0.061 |

Note: Calculated from Table 2

We are now ready to present our simulation results that show how the more generous potential benefit duration affects the steady-state unemployment rate (Table 5). We proceed as follows. To get the effect of the benefit duration increase from 30 to 52 weeks, we utilize the entire sample (all age groups, both before- and after-policy sample). Using our regression results of Table 2 (columns 2 and 4), we estimate, for each observation, the inflow- and outflow-probability with and without benefit duration extended to 52 weeks. With these estimates, we can calculate the implied steady-state unemployment rate with and without extended benefit duration. Moreover, using these estimates we can decompose the estimated increase in the steady-state unemployment rate into an inflow- and an outflow-component applying the procedure described. For the PBD extension from 30 to 39 weeks, we proceed in an analogous way.

Table 5 reports the result from this simulation exercise. In the upper panel of Table 5 we show the effects of the PBD extension from 30 to 52 weeks. The exit rate from unemployment (first column) falls from 0.4012 before the policy change and to 0.3680 after the change, implying a 8.3 percent reduction in the unemployment outflow probability.[17] Similarly, the entry rate into unemployment (second column) increases from 0.0098 to 0.0142, which amounts to a 44.9% increase in the unemployment inflow probability. Taken together, these estimates imply an increase in the steady-state unemployment-population ratio from 2.38% before the policy change to 3.72% after the policy

[17]Note that this result is very much in line with our earlier results on the effects of PBD extensions in Austria (Lalive et al. 2006) suggesting that extending PBD from 30 to 52 weeks increases unemployment duration by 2.27 weeks which is about 12% of average unemployment duration.

change.[18] In other words, the unemployment rate increases by 1.34 percentage points or by about 56%.

The lower panel of Table 5 applies the same procedure the estimate the effects of the PBD extension from 30 to 39 weeks. While qualitatively all effects go in the same direction they are quantitatively much smaller. The outflow probability decreases from .4649 to .4491 (a 3.4% reduction); and the inflow probability increases from 0.0098 to 0.0121 (an 23.5% increase).[19] These effects imply an increase in the equilibrium unemployment population ratio of 0.55 percentage points (or 26.6%), from 2.07% before the change to 2.62% after the change.

Table 5 shows a further interesting result. Decomposing the increase in the unemployment ratio into an inflow- and an outflow-component reveals that the bulk of the increase is due to the larger unemployment inflow rate. The effect of extended PBD on the unemployment outflow is much smaller. For the PBD increase to 52 weeks, 77.8% of the increase in the unemployment ratio can be attributed to an increase in the entry rate, whereas only 15.6% is due to a lower exit rate from unemployment. For the PBD increase from 30 to 39 weeks, an even larger fraction of the increase in the unemployment ratio (84.1%) is due to the increase in the inflow-rate, whereas only 13.1% can be assigned to the lower unemployment exit probability.

A further interesting indicator shows that the increase in PBD raises the unemployment ratio by 0.061 percentage points per additional PBD week for the extension from 30 to 52 weeks; and by 0.061 percentage points for the extension from 30 to 39 weeks. Interestingly, for both policy changes, the estimated effect per additional PBD week attributable to the unemployment inflow, is similar for the short (30 to 39) and the long (30 to 52) PBD increase. The isolated effect of one additional PBD week on the unemployment inflow indicates an increase in the unemployment ratio by 0.047 percentage points (increase from 30 to 52 weeks) and by 0.051 percentage points (increase from 30 to 39 weeks). The effects on the unemployment outflow are much smaller. We find that one additional PBD-week increases the equilibrium unemployment ratio by 0.010 percentage points for the policy change from 30 to 52 weeks, whereas the corresponding estimate for the policy change from 30 to 39 weeks amounts to only 0.008 percentage points.

### 6.1 Simulations for subgroups

We find that the increase in maximum benefit duration increases equilibrium unemployment, to some extent because the outflow from unemployment goes

[18] As indicated before, our sample contains attached workers for which the unemployment rate is rather low. For example, in the third quarter of 1988 the average unemployment rate in our sample was 2.04%.

[19] Note that the outflow result is, again, very much in line with our earlier result for Austria (Lalive et al. 2006) suggesting that extending PBD from 30 to 39 weeks increases unemployment duration by 0.45 weeks which is about 2% of average unemployment duration.

**Table 6** Decomposing the increase in the unemployment population ratio, various subsamples

| Subsample | u | Change in u | Due to outflow | | Due to inflow | | Due to interaction | |
|---|---|---|---|---|---|---|---|---|
| | | | Absolute | % | Absolute | % | Absolute | % |
| PBD 30 to 52 | | | | | | | | |
| Full sample | 0.0238 | 0.0134 | 0.0021 | 15.6% | 0.0104 | 77.8% | 0.0009 | 6.6% |
| Females | 0.0293 | 0.0198 | 0.0019 | 9.5% | 0.0169 | 85.3% | 0.0010 | 5.2% |
| Males | 0.0202 | 0.0062 | −0.0006 | −8.9% | 0.0069 | 111.8% | −0.0002 | −3.0% |
| Blue collar | 0.0234 | 0.0138 | 0.0002 | 1.7% | 0.0135 | 97.4% | 0.0001 | 0.9% |
| White collar | 0.0196 | 0.0111 | −0.0010 | −9.2% | 0.0128 | 115.1% | −0.0006 | −5.8% |
| Low wage | 0.0351 | 0.0170 | 0.0009 | 5.4% | 0.0157 | 92.3% | 0.0004 | 2.3% |
| High wage | 0.0168 | 0.0076 | 0.0001 | 0.8% | 0.0075 | 98.9% | 0.0000 | 0.3% |
| PBD 30 to 39 | | | | | | | | |
| Full sample | 0.0207 | 0.0055 | 0.0007 | 13.1% | 0.0046 | 84.1% | 0.0002 | 2.8% |
| Females | 0.0218 | 0.0034 | −0.0003 | −10.0% | 0.0038 | 111.8% | −0.0001 | −1.7% |
| Males | 0.0186 | 0.0044 | −0.0006 | −14.2% | 0.0052 | 118.1% | −0.0002 | −3.9% |
| Blue collar | 0.0210 | 0.0061 | −0.0004 | −6.0% | 0.0066 | 107.9% | −0.0001 | −1.8% |
| White collar | 0.0171 | 0.0025 | −0.0004 | −17.8% | 0.0030 | 120.9% | −0.0001 | −3.1% |
| Low wage | 0.0304 | 0.0050 | −0.0009 | −18.5% | 0.0061 | 122.0% | −0.0002 | −3.6% |
| High wage | 0.0148 | 0.0036 | 0.0005 | 13.6% | 0.0030 | 83.6% | 0.0001 | 2.7% |

Note: Calculated from Table 4

down but mainly through an increase in the inflow into unemployment. To investigate whether this results also holds for subgroups we use the parameter estimates presented in Table 4 to perform similar simulations as before, but now separately for each subgroup. Table 6 presents the simulation results. The upper part presents the results for the PBD change from 30 to 52 weeks, the lower part gives the simulation results for the PBD change from 30 to 39 weeks. For reasons of comparison the first rows of each part of the table replicates the main results of Table 5.[20]

As shown the PBD change from 30 to 52 weeks increases equilibrium unemployment for every subgroup with the increase for women, low wage workers and non-seasonal workers being larger than for their counterparts. There is hardly any difference between blue collar and white collar workers and between workers with low tenure and high tenure. For every subgroup the contribution to the change in equilibrium unemployment of the change in inflow is larger than that of the change in outflow.

Also for the PBD change from 30 to 39 weeks we find that the increase in equilibrium unemployment is mostly due to the increase in the inflow into unemployment and to a much smaller extent due to the decrease in the outflow from unemployment.

[20]Note that in the simulations we use all estimated parameters of Table 6 irrespective of whether or not they are significantly different from zero at conventional levels of significance.

## 7 Conclusion

According to job search theory an increase in the maximum unemployment benefit duration affects the unemployment rate both through a decrease in the outflow from unemployment and through an increase in the inflow to unemployment. These theoretical predictions are confirmed by empirical research. However, empirical research has been on either the outflow from unemployment or the inflow into unemployment. There are no studies that investigate both effects simultaneously. So, it is not clear to what extent effects on inflow and outflow affect the unemployment rate.

This paper uses microdata to evaluate the impact of an increase in maximum benefit duration on the steady-state unemployment rate distinguishing between these two effects. We draw on policy changes in Austria that extended maximum benefit duration from 30 to 52 (30 to 39) weeks for individuals above age 50 (between ages 40 and 49) with a continuous work history. We find that this policy change lead to 56% increase in the steady-state unemployment rate for the older age group and a 26% increase in the steady-state unemployment rate for the younger age group. Surprisingly, most of the increase in equilibrium unemployment is due to an increase in the inflow into unemployment, whereas the effect of the decrease in the outflow from unemployment is modest. We also find that the effects are stronger for women than for men. There may be institutional reasons for this as conditional on age women are closer to (early) retirement, and it is in line with the general notion that women react more strongly to incentives—wage elasticities of labor supply are larger for women than for men. Otherwise our results are rather robust across population subgroups.

In Lalive et al. (2006) we showed that the increase in PBD reduced the outflow from unemployment a lot. The PBD extension made it a lot more attractive for unemployed workers to reduce the search activities and thus lower their job finding rate. The novel findings in this paper are twofold. First, the extension of the PBD also increases the inflow into unemployment. The PBD extension made it only a little bit more attractive for employed workers to become unemployed. Our second finding is that the effect on the equilibrium unemployment rate due to increase in the inflow into unemployment is more important than the effect due to the decrease in the outflow from unemployment. Although the PBD extension makes it only a little bit more attractive for employed workers to become unemployed, there are many more employed workers than unemployed workers. It is the sheer mass of employed workers that cause the inflow effect to be larger.

From a policy point of view it is important to know that the inflow effect is larger than the outflow effect. Should this not be taken into account the effects of a change in PBD will be seriously underestimated. The fact that changes in PBD have quite a large—aggregate—inflow effect also means that PBD could be an instrument to increase the employment rate. If the PBD

is shortened, firms may become more reluctant to destroy jobs and it may also be less attractive for workers to "quit" into unemployment. We also note that our results are partly based on older (50+) workers which have low employment rates in many countries. Taking into account inflow effects for these groups seems highly relevant from a policy perspective. With respect to 50+ workers, it should be emphasized that our analysis has focused on flows between employment and unemployment only. However, often older workers who lost their job might consider to enter other programs (such as disability and early retirement) and changing unemployment insurance rules may affect flows into other programs. While this issue is beyond the scope of the present paper, future research should consider how changing incentives in one program may affect flows in and out of other programs.

**Acknowledgements** We thank the editor, Christian Dustmann, and three anonymous referees for helpful comments and suggestions. Andreas Steinhauer and Oliver Ruf did excellent research assistance. Financial support by the Austrian Science Fund (National Research Network S103: "The Austrian Center for Labor Economics and the Analysis of the Welfare State") is gratefully acknowledged.

## References

Adamchik V (1999) The effect of unemployment benefits on the probability of re-employment in Poland. Oxf Bull Econ Stat 61(1):95–108

Addison JT, Portugal P (2004) How does the unemployment insurance system shape the time profile of jobless duration? Econ Lett 85(2):299–234

Anderson PM, Meyer BD (1997) Unemployment insurance takeup rates and the after-tax value of benefits. Q J Econ 112(3):913–937

Card DE, Levine PB (2000) Extended benefits and the duration of UI spells: evidence from the New Jersey extended benefit program. J Public Econ 78(1–2):107–138

Carling K, Edin P-A, Harkman A, Holmlund B (1996) Unemployment duration, unemployment benefits, and labor market programs in Sweden. J Public Econ 59(3):313–334

Christofides LM, McKenna CJ (1995) Unemployment insurance and moral hazard in employment. Econ Lett 49(2):205–210

Christofides LM, McKenna CJ (1996) Unemployment insurance and job duration in Canada. J Labor Econ 14(2):286–312

Fredriksson P, Holmlund B (2003) Improving incentives in unemployment insurance: a review of recent research, IFAU working paper No.5

Fredriksson P, Holmlund B (2006) Improving incentives in unemployment insurance: a review of recent research. J Econ Surv 20(3):357–386

Green DA, Riddell WC (1997) Qualifying for unemployment insurance: an empirical analysis. Econ J 107(440):67–84

Green DA, Sargent TC (1998) Unemployment insurance and job durations: seasonal and non-seasonal jobs. Can J Econ 31(2):247–278

Grossman JB (1989) The work disincentive effect of extended unemployment compensation: recent evidence. Rev Econ Stat 71(1):159–164

Ham J, Rea S (1987) Unemployment insurance and male unemployment duration in Canada. J Labor Econ 5:325–353

Hunt J (1995) The effect of unemployment compensation on unemployment duration in Germany. J Labor Econ 13(1):88–120

Katz L, Meyer B (1990) The impact of the potential duration of unemployment benefits on the duration of unemployment. J Public Econ 41(1):45–72

Lalive R, Zweimüller J (2004a) Benefit entitlement and the labor market: evidence from a large-scale policy change. In: Agell J, Keene M, Weichenrieder A (eds) Labor market institutions and public policy 2003. MIT, Cambridge, pp 63–100
Lalive R, Zweimüller J (2004b) Benefit entitlement and unemployment duration: the role of policy endogeneity. J Public Econ 88(12):2587–2616
Lalive R, van Ours JC, Zweimüller J (2006) How changes in financial incentives affect the duration of unemployment. Rev Econ Stud 73(4):1009–1038
Layard R, Nickell S (1999) Labor market institutions and economic performance. In: Ashenfelter O, Card D (eds) Handbook of labor economics, vol. 3C, pp 3029–3084
Meyer BD (1990) Unemployment insurance and unemployment spells. Econometrica 58(4):757–782
Moffitt R (1985) Unemployment insurance and the distribution of unemployment spells. J Econom 28(1):85–101
Moffitt R, Nicholson W (1982) The effect of unemployment insurance on unemployment: the case of federal supplemental benefits. Rev Econ Stat 64(1):1–11
Mortensen DT (1977) Unemployment insurance and job search decisions. Ind Labor Relat Rev 30(4):505–517
Mortensen DT, Pissarides CA (1994) Job creation and job destruction in the theory of unemployment job creation and job destruction in the theory of unemployment. Rev Econ Stud 61(3):397–415
Nickell S, Layard R (1999) Labor market institutions and economic Performance. In: Ashenfelter O, Card D (eds) Handbook of labor economics. Elsevier, Amsterdam, pp 3029–3084
OECD (1999) Benefit systems and work incentives, Paris
Pissarides CA (2000) Equilibrium unemployment theory, 2nd edn. MIT, Cambridge
Puhani PA (2000) Poland on the dole: the effect of reducing the unemployment benefit entitlement period during transition. J Popul Econ 13(1):35–44
Roed K, Zhang T (2003) Does unemployment compensation affect unemployment duration? Econ J 113(484):190–206
Van Ours JC, Vodopivec M (2006) How shortening the potential duration of unemployment benefits entitlement affects the duration of unemployment: evidence from a natural experiment. J Labor Econ 24(2):351–378
Winter-Ebmer R (1998) Potential unemployment benefit duration and spell length: lessons from a quasi-experiment in Austria. Oxf Bull Econ Stat 60(1):33–45
Winter-Ebmer R (2003) Benefit duration and unemployment entry: a quasi-experimental evidence for Austria. Eur Econ Rev 47(2):259–273